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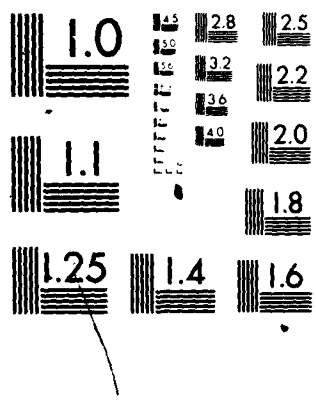

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Naval Studies Group

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PREDICTION WITH POOLED CROSS-SECTION AND TIME-SERIES DATA: TWO CASE STUDIES

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When estimating models with pooled cross-section and time-series data (e.g. estimating demand equations for all 50 states) one has to decide whether or not to pool the data. The usual procedure is to first test for the overall homogeneity (equality) of the coefficients. If this hypothesis is not rejected, then a single equation is estimated with pooled data. If the hypothesis is rejected, further hypothesis testing may be necessary. For example, if the model contains more than one coefficient the equality constraint may be rejected for only a subset of the coefficients. In this case the data is pooled and dummy variables are used with the subset of coefficients for which the equality constraint does not hold. There are at least three problems with this procedure of pooling (or not pooling) after some preliminary tests of significance. First, as noted in Maddala (~~77;~~ pp. 332-333) it raises problems about inference from the pooled model. Second, there is the related question of what significance level to use when deciding whether or not to pool. Third, the choice of estimates to select from is quite limited. That is, one must pick either the pooled or the non-pooled estimate, even if these two estimates are very different. These problems suggest that an alternative (or hybrid) method of handling pooled cross-section and time-series data is needed. The purpose of this paper is to propose such a method.

In Section I the hybrid procedure is presented. In Section II the predictive power of the hybrid estimates are compared with the predictive power of the pooled and non-pooled estimates. Section III contains the conclusions.

I. The Hybrid Estimation Technique

Consider the simple model:

(1) $X_{it}=a_i+b_iX_{it}+U_{it}$; $i=1,\ldots,N$; $t=1,\ldots,T_i$,

where $U_{it} \sim N(0,\sigma_i^2)$. The i subscript represents N cross-sections and the t subscript represents T_i observations for the i^{th} cross-section. Since equation(1) is to be estimated with pooled cross-section and time-series data, three general approaches to the estimation (and prediction) problem are possible.

In the first approach, ordinary least squares is applied to each of the N cross-sections separately. That is, the OLS *estimates* a_i and b_i are calculated from:

(2) $$\hat{b}_i = \frac{S_{X_iY_i}}{S_{X_iX_i}},$$

(3) $$\hat{a}_i = \bar{Y}_i - \hat{b}_i\bar{X}_i,$$

where

(4) $$S_{X_iX_i} = \sum_{t=1}^{T_i} (X_{it} - \bar{X}_i)^2,$$

(5) $$S_{X_iY_i} = \sum_{t=1}^{T_i} (X_{it} - \bar{X}_i)(Y_{it} - \bar{Y}_i).$$

However, this approach seems quite restrictive. The estimates for each individual cross-section are based only on the information contained in the sample for that particular cross-section. This implicitly assumes each cross-section is so different that nothing can be gained by observing analogous behavior in other cross-sections. A more reasonable assumption might be that useful information about behavior in any one cross-section is contained in the behavior of other cross-sections.

In the second approach, the data are pooled. That is, the pooled estimates $\hat{a}_p$ and $\hat{b}_p$ are calculated from:

(6) $$\hat{b}_p = \frac{T_{xy}}{T_{xx}}$$

(7) $$\hat{a}_p = \bar{Y} - \hat{b}_p\bar{X},$$

where

(8) $$T_{xx} = \sum_{i,t} (X_{it} - \bar{X})^2,$$

(9) $$T_{xy} = \sum_{i,t} (X_{it} - \bar{X})(Y_{iy} - \bar{Y}).$$

This approach has the advantage of using all the information in the sampel to estimate each cross-section's coefficients. The disadvantage is that the coefficients for each cross-section are constrained to equality. This seems quite restrictive in that all cross-sections are assumed to be exactly alike with respect to the behavior represented in equation 1.

The usual procedure is to choose between the first and second approach on the basis of an F-test.[1] However, both of these choices seem extreme. In addition, it is not clear which approach will provide better predictions. All this suggests that a method which combines the advantages of the first two approaches might be a better alternative. The third approach to the estimation and prediction problem is thus a hybrid of the first two approaches.[2] In this third approach, a time series equation is first estimated separately for each cross-section, and then these coefficients are adjusted toward the common coefficient *from* the *pooled* regression. The purpose of this approach is to use all the available information in the estimation of the coefficients for each cross-section, but at the same time to allow for [illegible] differences among individual cross-sec[illegible]

The hybrid method can be briefly described as follows. Assume that there are two pieces of information about the true parameter b_i, where b_i is a coefficient to be estimated for the i^{th} cross-section. The first piece of information $\hat{b}_i$, where $\hat{b}_i$ is the ordinary least squares estimate of b_i based only on the observations for the i^{th} cross-section. The second piece of information is $\hat{b}_p$, where $\hat{b}_p$ is the estimated coefficient when the data are pooled over all cross-sections. According to the hybrid approach, a reasonable estimate of b_i is a weighted average of $\hat{b}_i$ and $\hat{b}_p$. More specifically, the hybrid estimates b_i^* and a_i^* are defined as:

$$(10)\quad b_i^* = \frac{W_1\hat{b}_i + W_2\hat{b}_p}{W_1 + W_2},$$

$$(11)\quad a_i^* = \bar{Y}_i - b_i^*\,\bar{X}_i,$$

where

$$(12)\quad W_1 = \left(\frac{S_{x_ix_i}}{\hat{\sigma}_i}\right)\left(\frac{S_{y_iy_i}/(T_i-1)}{\frac{1}{N}\sum_{i=1}^{N}\left[S_{y_iy_i}/(T_i-1)\right]}\right),$$

$$(13)\quad W_2 = \left(\frac{N}{N-1}\right)\left(\frac{1}{Z}\right),$$

$$(14)^{3}\quad Z = \frac{1}{N-1}\sum_{i=1}^{N}(b_i-b_p)^2 - \frac{1}{N}\sum_{i=1}^{N}\frac{\hat{\sigma}_i^{\,2}}{S_{x_ix_i}},$$

$$(15)\quad S_{y_iy_i} = \sum_{i=1}^{T_i}(Y_i - \bar{Y}_i)^2.$$

If the model in equation 1 has K (K > 1) independent variables the analysis is similar, but the term $\frac{\hat{\sigma}_i^{\,2}}{S_{x_ix_i}}$ is replaced by the estimated variance of the OLS coefficients ($\hat{V}_{ik}$; i=1,...,N; k=1,...K), where V_{ik}'s are obtained from the N cross-section regressions. If only a subset (K_s) of the K coefficients are assumed heterogeneous, then the vectors $\bar{b}_i$ are estimated from a pooled regression, but dummy variables are used on the heterogeneous subset $\bar{b}_{s_i}$. In this case the term

$\frac{\hat{\sigma}_i^{\,2}}{S_{x_ix_i}}$ is replaced by $\frac{\hat{\sigma}_i^{\,2}}{\hat{\sigma}^2}\hat{V}_{ik}$, i=1,...,N; k=1, ... K_s; where:

$$(16)\quad \hat{\sigma}^2 = \frac{\sum_{i,t}(Y_{it} - X_{it}\,\hat{\bar{b}}_i)^2}{df_p}$$

$$(17)\quad \hat{\sigma}_i^{\,2} = \frac{\sum_{i=1}^{T_i}(Y_i - X_i\,\hat{\bar{b}}_i)^2}{T_i - K - 1};\ i=1,\ \dots\ N,$$

(18) df_p = degrees of freedom in the pooled with dummies regression. All other terms, including $\frac{S_{x_ix_i}}{\hat{\sigma}_i}$, are the dame when the model has more than one independent variable.[4]

In the single variable model, W_1 depends on three terms: (1) the standard error of the regression of the i^{th} cross-section; (2) the variance of the dependent variable in the i^{th} cross-section relative to the average variance for all cross-sections; and (3) the sum of squared deviations of the independent variable for the i^{th} cross-section. As the standard error of the regression for the i^{th} cross-section is smaller, W_1 is larger. As the relative variance of the dependent variable increases, W_1 increases. W_1 is also greater, the greater is the sum of square deviations of the independent variable. This method of determining W_1 is intuitively appealing. The better the fit for an individual cross-section, as measured by a smaller stnadard error of the regression, the closer b_i^* is to $\hat{b}_i$. However, the standard error of the regression depends on the variance of the dependent variable. To correct for the fact that the standard error of the regression may be small (large) just because the variance of the dependent variable is small (large), the relative variance of the dependent variable is also included in W_1. Finally, as the variance of the independent variable increases (or as the number of observations increases), more information is contained in the estimate of $\hat{b}_i$. Therefore, as the sum of squared deviations of the independent variable increases, more weight is placed on $\hat{b}_i$.[5]

W_2 depends on the variability of the $\hat{b}_i$'s around $\hat{b}_p$ relative to the average variance of the $\hat{b}_i$'s. As this relative variance increases, W_2 becomes smaller and less weight is placed on $\hat{b}_p$. This is also intuitively appealing. As the variability of the $\hat{b}_i$'s around $\hat{b}_p$ increases, there is less reason to think that $\hat{b}_p$ contains useful information about the individual b_i. On the other hand, as the average variance of the estimated $\hat{b}_i$'s increases, there is less reason to think that each $\hat{b}_i$ contains useful information about the individual b_i.

In summary, the hybrid procedure adjusts each $\hat{b}_i$ toward $\hat{b}_p$, but puts more weight on $\hat{b}_i$ as $S_{x_ix_i}$ is larger, $\hat{\sigma}_i$ is smaller, $\frac{S_{y_iy_i}/(T_i-1)}{\frac{1}{N}\sum_{i=1}^{N}\left[S_{y_iy_i}/(T_i-1)\right]}$ and Z is larger. When the model contains more than one independent variable the procedure is similar except for the slight differences noted earlier.[6]

II. An Application: Two Case Studies

In this section the hybrid procedure is applied to two different case studies. The two cases to be analyzed arise in entirely different contexts, but both involve the use of pooled cross-section and time-series data to investigate current policy problems. The first case is the prediction of inflation in fifteen Latin American countries over the period 1973-1975, based on coefficients estimated from the 1954-1972 period. The second case is the prediction of receipts for 49 state governments for fiscal year 1975 based on coefficients from the 1958-1974 period. In both cases predictions based on pooled regressions are compared with predictions based on individual (country or state) regressions and with predictions based on the adjusted coefficients as suggested by the hybrid method.

Case 1: Inflation in Latin America

The equation to be estimated in the first case study was initially formulated by Harberger (1963) in his study of Chilean inflation and subsequently estimated by Vogel (1974) for sixteen Latin American countries. The equation is:

$$(19)\quad P_t' = b_o + b_1 M_t' + b_2 M_{t-1}' + b_3 Y_t' + b_4 (P_{t-1}' - P_{t-2}') + u_t$$

where M is the money supply, Y is real GDP, P is the consumer price index, primes indicate annual percentage changes, subscripts indicate current or lagged variables, u is the error term, and b_0 through b_4 are the coefficients to be estimated.

This equation has been estimated for fifteen Latin American countries over the period 1954 through 1972.[7] Based on the results from the individual country regressions and the pooled regression, the F-test indicates that the data should not be pooled. A similar conclusion was reached by Vogel (1974). However, it is not clear which approach will provide better predictions. Moreover, Vogel (1974) emphasizes that allowing all coefficients to differ among countries (i.e. not pooling) only improves the $\bar{R}^2$ by .03 over the $\bar{R}^2$ obtained in the pooled regression. For the countries and time period covered in the present study the improvement in $\bar{R}^2$ is about .06. These findings suggest that heterogeneity among the countries is quite limited -- a surprising result in light of the widely differing inflationary experiences of Latin American countries and the importance frequently attributed to the structural factors excluded from the model.

The preceeding discussion suggests that comparing the predictive power of the pooled, non-pooled and hybrid estimates would be a useful exercise. These predictions are presented in Table 1. The predictions are based on coefficients estimated over the period 1954 through 1972, and the rate of inflation has been predicted for each of the fifteen Latin American countries for the period 1973 through 1975. For each set of predictions, three types of statistics are given: mean error, mean absolute error, and mean square error. These statistics are reported for each country individually, for the aggregate of all fifteen Latin American countries, and for the aggregate excluding Chile. Chile completely dominates the aggregate results because of its astronomical rate of inflation which averaged more than 400 percent per year over the 1973-1975 period. Hence it is useful to examine the aggregate results excluding Chile as well as the results for each individual country.[8]

The coefficients from the pooled regression provide the best predictions in the aggregate including Chile, but only because the underprecition of Chilean inflation is slightly less serious with this method than with the alternative methods. When Chile is excluded, all three methods give much better predictions in the aggregate. In particular, the coefficients from the pooled regression underpredict the rate of inflation by only a little more than 1 percent on the average, while the hybrid method underpredicts by just over 2 percent. These results seem quite good considering that average rates of inflation ranged from approximately 8 percent in Honduras and Venezuela to more than 80 percent in Argentina and Uruguay for the 1973-1975 period. The coefficients from the individual country regressions yield the predictions with the lowest mean square error in the aggregate excluding Chile, but the mean square error for the hybrid method is only slightly higher. Overall, the hybrid method appears to give the best predictions in the aggregate excluding Chile. Not only is the hybrid method a very close second in terms of the mean error and mean square error criteria, but it also produces the lowest mean absolute error.

Turning to the results for the individual countries, none of the methods clearly predominates in predicting inflation in the five Latin American countries which historically have experienced high rates of inflation. As indicated above, the coefficients from the pooled regression yield the best predictions for Chile, and the same is true for Brazil. On the other hand, the coefficients from the individual regressions yield the best predictions for Argentina, Bolivia, and Uruguay. However, for Bolivia the hybrid method gives equally good predictions, and for Uruguay only one observation is available for comparison.[9]

For the ten Latin American countries with moderate or low rates of inflation, the results are much more clearcut. For all of these countries except Venezuela the coefficients from the pooled regression or the hybrid method give the best predictions of inflation. Moreover, the coefficients from the individual regressions give the worst predictions for eight of these ten countries.[10] It is thus apparent that useful information about the behavior of inflation in individual Latin American countries, especially moderate and low inflation countries, can be obtained by examining the behavior of inflation in the other Latin American countries.

Case 2: A Model of State Government Receipts

In the second case study, the total receipts of forty-nine state governments are analyzed. The equation to be estimated is:

(20) $\text{Total Receipts}_t = b_1 \text{ Potential GNP}_t + b_2 \text{ Real Income}_t + b_3 \text{ Deflator}_t + b_4 \text{ Surplus}_{t-1} + u_t$

where all variables are measured in percentage changes, except the surplus which is the level in the preceding year; u_t is the error term, and b_1 through b_4 are the coefficients to be estimated. The coefficient of potential GNP indicates the long-run elasticity of total receipts with respect to the growth in potential output. The most important coefficient, b_2, indicates the elasticity of total receipts with respect to short-run fluctuations in real income. These fluctuations in economic activity are measured by state real personal income (where state personal income is deflated by the implicit price deflator for GNP). The elasticity of total receipts with respect to changes in the price level is indicated by the coefficient of the deflator, which is measured by the implicit price deflator for state and local purchases of goods and services. Finally, the level of the surplus (or deficit) in the preceding year is measured for each state by local receipts less total expenditures as a percentage of total expenditures. This variable is introduced because the greater the level of the surplus (deficit) in the preceding year, the less (more) will be the need for tax increases during the ensuing year, and hence the smaller (larger) will be the percentage increase in total receipts.

The statistical analysis which follows covers only 49 of the 50 state governments. Alaska has been excluded because of very large and erratic grants from the Federal Government. Data on the individual states are from the Bureau of the Census, State Government Finances, for the fiscal year 1957 through 1975. All states but three have fiscal years which begin on July 1 and end on June 30, and for the independent variables quarterly averages to conform to these fiscal years have been used. Total receipts of each state are measured net of receipts of state unemployment compensation funds and net of receipts of state social insurance funds (e.g., public employee retirement systems). These funds have been excluded because they are handled separately and cannot be mingled with general state funds.

All variables have been expressed in percentages for a variety of reasons, but the most important reason is so that the data can readily be pooled across states. In addition, when the variables are expressed as percentage changes, the estimated coefficients are elasticities which are constant for all values of the dependent and independent variables.

To test for the homogeneity of all coefficients in equation 20, the F-test for reduction in residual sums of squares is 1.33 (with degrees of freedom 192,637). Based on the interpolation of the F-tables, this is significant at the 1 percent level.

To explore further this apparent heterogeneity among states, F-ratios have been calculated for the individual coefficients. When the coefficients for state real personal income are allowed to differ among states, but the remaining coefficients are required to be the same across states, the F-ratio for reduction in residual sums of squares is 1.73 (with degrees of freedom 48, 781). This is significant at the 1 percent level by a comfortable margin. Thus, the response of receipts to fluctuations in economic activity does differ significantly among states when economic fluctutations are measured by state real personal income. When the test for significant reduction in residual sums of squares is applied to the remaining coefficients, given that the coefficients of state real personal income differ among states, the F-ratio is 1.18 (with degrees of freedom 144,685). This is not significant at the 5 percent level, so the response of receipts to potential GNP, the deflator, and the surplus are assumed to be homogeneous across states.

To test the predictive ability of the pooled, non-pooled and hybrid estimates, the percentage change in receipts for each of the 49 states has been predicted for fiscal 1975. These predictions are presented in Table 2. They are based on the coefficients estimated over the period fiscal 1958 through fiscal 1974.[11] For each set of predictions, three statistics are given: mean error, mean square error and mean absolute error. The results are clearcut. All three approaches do equally well on mean absolute error. The pooled coefficients underpredict the percentage increase in state receipts by somewhat less than the individual state coefficients for state real personal income, but the pooled coefficients produce a slightly larger mean square error. However, the hybrid method yields both the smallest underprediction and the lowest mean square error. These results suggest that the coefficients of state real personal income produced by the hybrid method provide a better indication of the responsiveness of state government receipts to fluctuations in economic activity than either the pooled or individually estimated coefficients.

III. Conclusions

In this paper we have proposed a hybrid estimation technique that is useful when analyzing pooled cross-section and time-series data. The hybrid estimates are a weighted combination of the pooled and individual regression coefficients. The idea of this approach is to use all the available information for the estimation of each regression parameter, while at the same time allowing for potential differences that may exist between the different cross-sections. To test this method, we compared its predictive ability to the predictive ability of both the pooled and individual regression coefficients. This predictive test was performed on two separate models: a model of inflation in fifteen Latin American countries, and a model of dollar receipts for 49 state governments. In both instances the hybrid estimates yielded better prediction than either the pooled or individual regression estimates.

FOOTNOTES

1. The appropriate F-tests for pooling are well known; see, for example, Maddala (1971).

2. An approach similar to the one used in this paper is discussed in Maddala (1977; pp. 332-333).

3. A negative Z implies that W_2 is very large. Consequently, b_i^* should simply be the pooled estimate.

4. Of course, if only a subset of the coefficients are assumed heterogeneous, $\hat{\sigma}_i^2$ is as defined in (17). Also, if no intercept term appears in the model, the degrees of freedom in (17) are T_i-K.

5. Consider, for example, a simple model where inflation is a function of money supply in each of several countries. If in a particular country there is little variation in the money supply, little information can be obtained about the effect of money on prices, and hence relatively little weight should be placed on $\hat{b}_i$. On the other hand, as the variance or the number of observations increases, more information is provided about money and prices in that particular country, so more weight should be placed on $\hat{b}_i$.

6. With more than one explanatory variable each $\hat{b}_{ik}$ is adjusted toward $\hat{b}_{pk}$ independently of the remaining K-1 coefficients. That is, the covariance terms $S_{x_k x_j}$ ($k \neq j$) are not taken in the b_{ik}^* calculation.

7. The data are taken from the International Monetary Fund, International Financial Statistics (Washington, D. C.: May 1977). The regressions begin with the observations for 1954, rather than 1952, because of the presence of lagged variables in the equation. For Uruguay and Honduras the initial observations are respectively 1956 and 1957 because some of the earlier data are not available for these two countries.

8. While Chile's astronomical inflation can be adequately explained by monetary factors or by the economic chaos surrounding the military coup and Allende's death is debatable. In any case, it is not appropriate to evaluate the aggregate results solely on the ability (or inability) to predict inflation in only one of the fifteen countries.

9. For Uruguay the rate of inflation can be predicted only for 1973 because money supply data are not available for 1974 and 1975.

10. It may be that Venezuela's role as a major oil exporter makes the inflationary behavior of the other Latin American countries irrelevant. Moreover, the only other country in the low and moderate inflation group for which the coefficients from the individual regressions do not give the worst prediction is Ecuador, the only Latin American member of O.P.E.C.

11. To make the predictions of all three methods directly comparable, the coefficients from the pooled regression were always used for the deflator, potential GNP and surplus variables. This was done to remove all effects on the predictions that are not a direct result of the three different approaches used to estimate b_2.

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Table 1

MEAN SQUARE ERRORS FROM PREDICTIONS OF INFLATION FOR FIFTEEN LATIN AMERICAN COUNTRIES, 1973-1975**

	Individual Regressions	Pooled Regressions	Hybrid Method
	Mean Square Error	Mean Square Error	Mean Square Error
Aggregate	.2890	.2701	.4473
Aggregate excluding Chile	.0186	.0282	.0190
Uruguay*	.1192	.1491	.1279
Bolivia	.0278	.0403	.0280
Chile	3.8945	3.4946	6.1571
Brazil	.0182	.0060	.0179
Argentina	.0691	.1917	.1184
Paraguay	.0051	.0033	.0039
Colombia	.0076	.0045	.0056
Peru	.0048	.0044	.0007
Mexico	.0062	.0023	.0029
Ecuador	.0023	.0033	.0020
Costa Rica	.0298	.0037	.0112
Honduras	.0052	.0031	.0023
Venezuela	.0081	.0629	.0162
El Salvador	.0098	.0011	.0007
Guatemala	.0136	.0003	.0013

*For Uruguay the rate of inflation is predicted only for 1973 because necessary data are not available for 1974 and 1975.

**The countries are arranged from highest to lowest rates of inflation during the 1952-1972 period.

For lack of space the mean error and mean absolute error are not reported here.

Table 2

COMPARISON OF PREDICTIONS OF RECEIPTS FOR 49 STATE GOVERNMENTS, 1975

	Mean Error	Mean Square Error	Mean Absolute Error
Pooled Regression	.0100	.00381	.0486
Individual Coefficients for State Real Personal Income	.0132	.00360	.0478
Hybrid Method	.0094	.00359	.0481

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